Inflation targeting and private sector forecasts

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Abstract

Transparency is one of the biggest innovations in central bank policy of the past quarter century. Modern central bankers believe that they should be as clear about their objectives and actions as possible. However, is greater transparency always beneficial? Recent work suggests that when private agents have diverse sources of information, public information can cause them to overreact to the signals from the central bank, leading the economy to be too sensitive to common forecast errors. In this way, greater transparency could be destabilizing. While this theoretical result has clear intuitive appeal, it turns on a combination of assumptions and conditions, so it remains to be established that it is of empirical relevance.

In this paper we study the degree to which increased information about monetary policy might lead to individuals coordinating their forecasts. Specifically, we estimate a series of simple models to measure the impact of inflation targeting on the dispersion of private sector forecasts of inflation. Using a panel data set that includes 15 countries over 20 years we find no convincing evidence that adopting an inflation targeting regime leads to a reduction in the dispersion of private sector forecasts of inflation. While for some specifications adoption of inflation target does seem to reduce the standard deviation of inflation forecasts, the impact is rarely precise and always small.

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I. Introduction

Transparency is one of the biggest innovations in central bank policy of the past quarter century. Modern central bankers believe that they should be as clear about their objectives and actions as possible. This notion arises from the view that policymakers should be a source of stability, not a source of noise, with the economy and markets responding to data, not to the policymakers themselves.

Inflation targeting is one of the earliest and most comprehensive implementations derived from this view. As a framework for monetary policy, inflation targeting involves "the public announcement of medium-term numerical targets for inflation [and] increased transparency of the monetary policy strategy through communication with the public and the markets about plans, objectives, and decisions of the monetary authority" (Mishkin (2001)). The result is not only clearly understood and published numerical targets, but also inflation reports that explain past and likely future actions. But transparency is not nudity. Understanding policymakers' contingency plans does not mean laying the policymaking process bare for all to see. Monetary policymakers should not put cameras in the meeting room. There are clear limits. What are they?

Recent theoretical work has put this question into a new perspective. In their pioneering work, Morris and Shin (2002) show that when private agents have diverse sources of information, public information can cause them to overreact to the signals from the central bank leading the economy to be too sensitive to common forecast errors. The reason for this is that individuals care not only about accurately estimating the state of the economy, but also about having an estimate that is not too different from that of others. The implication is that more transparency may in fact be destabilising, so policymakers should think long and hard before changing their disclosure policies in ways that publicise more information.

Svensson (2006) and Woodford (2005) both suggest that the Morris and Shin result is likely to be a theoretical curiosum rather than anything policymakers should worry about. That is, the circumstance under which additional information is welfare reducing is extremely unlikely to occur in the real-world. As Svensson shows, Morris and Shin's own conclusion only holds when the noise in policymakers' publicly announced information is at least eight times that of the private information agents have obtained on their own. That is, public officials must be far worse in their evaluations of the economic environment that private agents. Evidence, such as that in Romer and Romer (2000), suggest that central bank staff forecasts are at least as good, if not better than, those of market economists.

Woodford's critique is based on the Morris and Shin choice of how to aggregate the quadratic loss functions of the individual agents. In their original paper, Morris and Shin assume that policymakers seek to minimise a social loss function that is based on the average squared error of individual estimates of the state of the economy. By contrast, if the social welfare function includes losses associated with the dispersion of agents' estimates of the state – something Morris and Shin assume the agents themselves care about – then more information is unambiguously a good thing.

Regardless of these two coherent and largely convincing criticisms, the Morris and Shin argument retains intuitive appeal. Policymakers worry that releasing more information might cause private agents to coordinate expectations, leaving the economy more exposed to common shocks. In the end, however, we are left with an empirical question: Does increased transparency lead to lower dispersion in private forecasts? If the answer is yes, we would need to look further; if the answer is no, then there is little to worry about.

In this paper we study the degree to which increased information about monetary policy might lead to individuals coordinating their forecasts. By combining information about whether a country targets inflation with the dispersion of private sector forecasts of inflation, we seek to measure the degree to which increased information leads individuals to coordinate their forecasts.

In order to examine this, we estimate a series of simple models designed to measure the impact of inflation targeting on the dispersion of private sector forecasts of inflation. Using a panel data set that includes 15 countries over 20 years we find no convincing evidence that adopting an inflation targeting regime leads to a reduction in the dispersion of private sector forecasts of inflation. While for some specifications adoption of inflation target does seem to reduce the standard deviation of inflation forecasts, in others it does not. And the precision of the estimates is rarely very high. The bulk of our evidence does not support the view that a shift to inflation targeting has resulted in a significant decline in the cross-sectional standard deviation of inflation forecasts across survey respondents.

Before proceeding, it is useful to note that our work is distinct from, but related to, two earlier papers. First, Mankiw, Reis, Wolfers (2004) examine the dispersion of inflation expectations in survey data and find that inflation expectations have become more concentrated around the mean as the level of inflation has fallen. At first glance this may seem as if it is a result that is more positive than ours. But, given that Mankiw, Reis and Wolfers only study US data, it is not possible to disentangle the impact of disinflation from increased Federal Reserve transparency.

Levin, Natalucci, and Piger (2004) investigate how well the mean of inflation expectations have been anchored, also from survey data. They provide evidence on how inflation targeting changed the dynamics of inflation. Their results suggest that the adoption of an explicit inflation target reduces the correlation of long-run inflation expectations with short-run movements in inflation, largely eliminating the link between expectations and realised inflation. Furthermore, Levin, Natalucci and Piger find that the adoption of an inflation targeting framework lowers the persistence of inflation, behaving more closely like a random walk.

The remainder of this paper is organised in five sections. Section II provides a description of the data we use. This is followed in section III with a simple statistical analysis, and in section IV with the results of more sophisticated regressions. Section V provides a conclusion.

II. Description of data

Beginning with the data, we study the dispersion of monthly survey-based inflation expectations for a number of countries from October 1989 to April 2009. The data are collected by Consensus Forecasts. Each month the firm surveys a large cross-section of professional forecasters – currently more than 7000 world-wide – asking each one for their current and next years' predictions for growth, inflation, unemployment, and short- and long-term interest rates in the countries that they follow. For each month, for each variable, Consensus Forecasts reports the high, low, and median forecast, as well as the standard deviation of survey responses. While Consensus Forecasts supplies forecast information for more than 70 countries, we restrict ourselves to the following 16: Canada, the Euro area, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, the UK, the US, Australia, and New Zealand. For many of the results, we ignore the Euro area because data are only available starting in December 2002.

This sample is sufficiently diverse to allow us to study the impact of inflation targeting, as two countries (New Zealand and Sweden) targeted inflation over the entire period, six (UK, Switzerland, Spain, Norway, Canada, and Australia.) adopted inflation targeting at some point during the sample, and the remaining eight have never adopted an explicit inflation target. For the second group, the six that adopted an inflation target during the 1990s, we need to choose a date for the adoption. It is perhaps surprising that there is disagreement on this timing. Mishkin and Schmidt-Hebbel (2007), Ball and Sheridan (2005), and Truman (2003), among others, all choose slightly different dates. For the most part, we adopt the dating in Appendix A of Mishkin and Schmidt-Hebbel (2007).

Table 1

Dating inflation regimes

	Inflation targeting regime	Non-inflation targeting regime
Australia	June 1993 to April 2009	November 1990 to May 1993
Canada	February 1991 to April 2009	October 1989 to January 1991
Euro Area		December 2002 to April 2009
France		October 1989 to April 2009
Germany		October 1989 to April 2009
Italy		October 1989 to April 2009
Japan		October 1989 to April 2009
Netherlands		January 1995 to April 2009
New Zealand	December 1994 to April 2009	
Norway	March 2001 to April 2009	June 1998 to February 2001
Spain	January 1995 to June 1998	July 1998 to April 2009
Sweden	January 1995 to April 2009	
Switzerland	January 2000 to April 2009	June 1998 to December 1999
United Kingdom	October 1992 to April 2009	October 1989 to September 1992
United States		October 1989 to April 2009

Notes: The dates shown correspond to dates for which the dispersion of inflation and GDP forecasts are available. They do not necessarily correspond to dates at which a country adopted an inflation target.

Source: Appendix A, Mishkin and Schmidt-Hebbel (2007), Norges Bank Regulation on Monetary Policy, March 29, 2001. Contrary to the date given in Mishkin and Schmidt-Hebbel, the start date of inflation targeting for Australia is June 1993, based on data from the Reserve Bank of Australia.

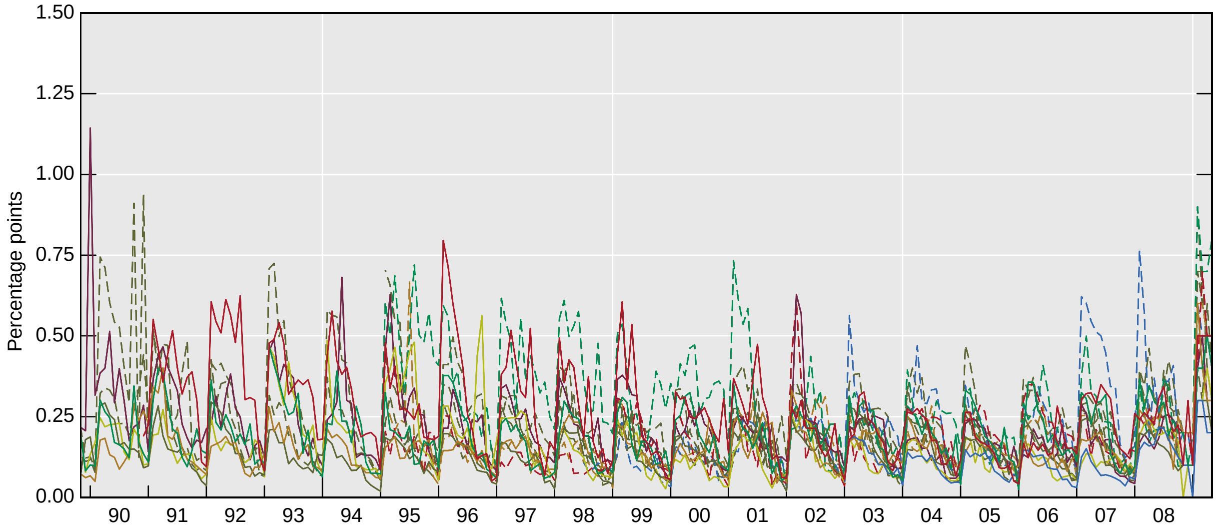
To continue, we need a bit of notation. We use the general form $S_{it}(.,.)$ to denote the standard deviation of private sector forecasts for country *i* made on date *t*. Next, we specify the variable being forecasted as π for CPI inflation and *y* for GDP growth, and whether the forecast is for the current year, which we denote by *c,* or for the next year, which we denote by *n*. Using this notation, $S_{it}(\pi, c)$ is the standard deviation of private sector forecasts for CPI inflation made at date *t* for the current year (the year containing *t*). Analogously, $S_{it}(\pi, n)$ is the standard deviation of private sector forecasts for CPI inflation also made at date *t* but for the next year (the year containing *t*+1), and $S_{it}(y, c)$ and $S_{it}(y, n)$ are the standard deviations of private sector forecasts for GDP growth for the current and next year, respectively.

To provide a sense of the time-series properties of the data, Graphs 1 and 2 plot the standard deviation of current and next year's forecasts of inflation, $S_{it}(\pi, c)$ and $S_{it}(\pi, n)$, for all the countries in our sample. Simple inspection reveals that the standard deviation of forecasts for the current year, $S_{it}(\pi, c)$ plotted in Graph 1, has significant seasonality; while $S_{it}(\pi, n)$ has less seasonality. Focusing on $S_{it}(\pi, c)$, a closer look shows that the standard deviation is highest in January and falls throughout the year. This is not at all surprising, since as a particular year progresses, inflation during that year increasingly becomes an historical fact that need not be estimated.

Graph 1

Standard deviation of current year's inflation forecasts, $S_{it}(\pi,c)$

Monthly data

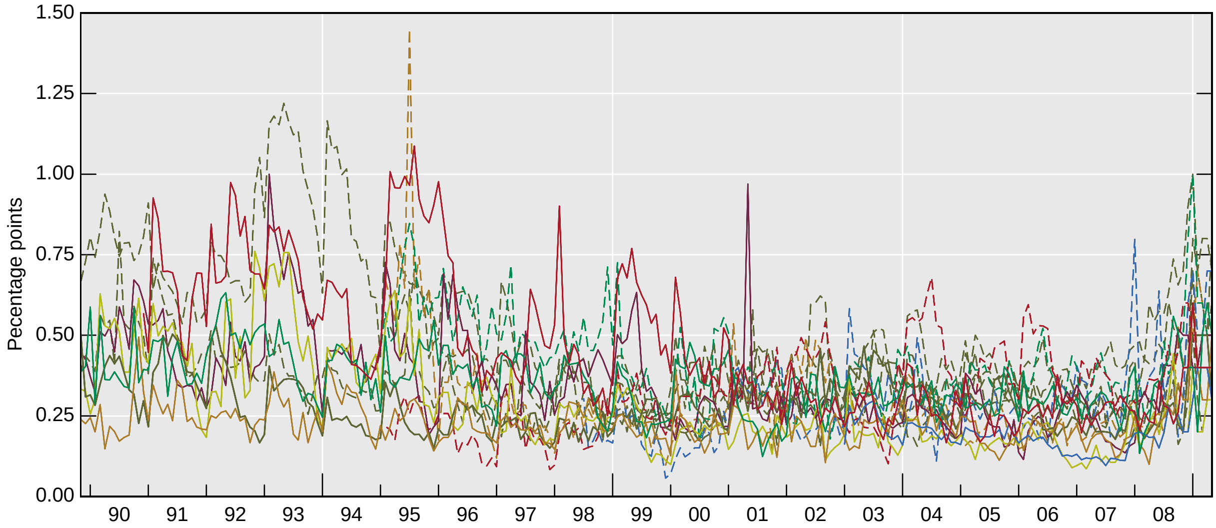


Source: Consensus Economics.

Graph 2

Standard deviation of next year's inflation forecasts, $S_{it}(\pi,n)$

Monthly data



Source: Consensus Economics.

To assess the seasonality in these series, we assume it is deterministic and estimate the following regressions:

(1) $$S_{it}(\pi,c) = \sum_{k=1}^{12} \beta_k^c D_{kt}^m + \varepsilon_{it}$$

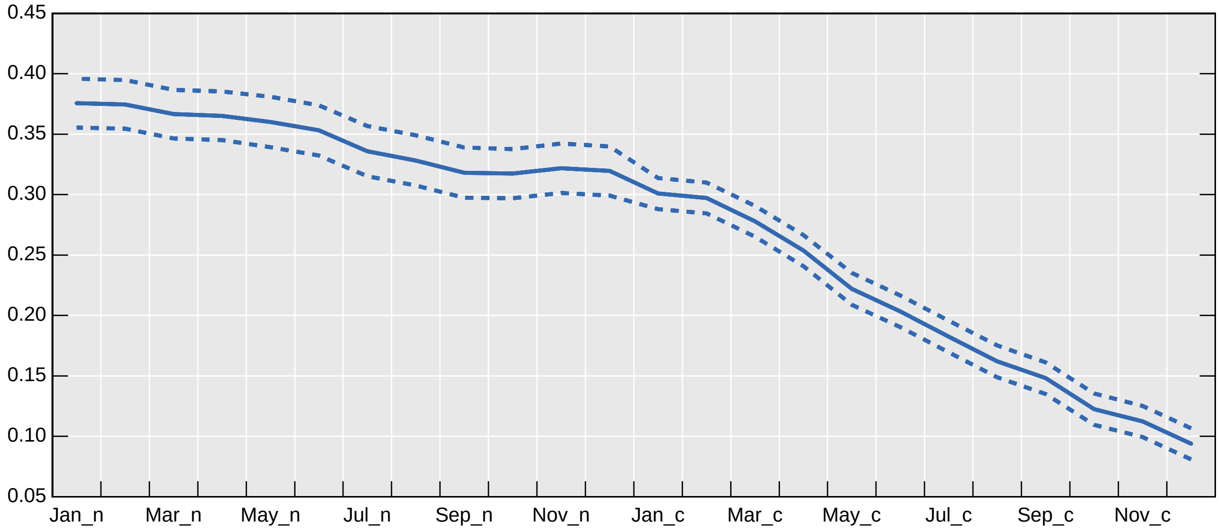
and

(2) $$S_{it}(\pi,n) = \sum_{k=1}^{12} \beta_k^n D_{kt}^m + \varepsilon_{it} \,,$$

where D_{kt}^m is a monthly dummy for month k. Notice that for this exercise, the coefficients (the β_k's) are constrained to be equal across all of the countries in the sample.

Graph 3

Seasonal dummy coefficients



Source: Author's calculations based on equations (1) and (2).

The coefficients from equations (1) and (2), with a 95% confidence interval, are plotted in Graph 3. We report the coefficients from left to right depending on the amount of time from the date of the survey to the end of the period covered by the forecast. Starting on the far left, the first observation, labelled "Jan_n," is the coefficient associated with the dummy variable for the standard deviation of the January forecast for the next year's inflation (β_1^n), which is completed 24 months in the future. By contrast, the far right of the figure plots the coefficient associated with the standard deviation of inflation in December of the current year (β_{12}^c), which is completed in only one month. While the relationship is not linear, it is clearly declining. And a regression of $\{\beta_1^n, \beta_2^n \ldots, \beta_{11}^c, \beta_{12}^c\}$ against {1, 2, ...,23, 24} yields a slope coefficient of -0.012 and a t-statistic of 16.2.[1] This means that the marginal effect of an additional month of data reduces the standard deviation of private inflation forecasts by -0.012. The average value of $S_{it}(\pi, n)$in January (the standard deviation of private sector forecasts for "next" year in January of "this" year) is 0.377 percentage points and the average value of $S_{it}(\pi, c)$ in December is 0.094 percentage points. Thus, each additional month of data tightens the spread of private sector inflation forecasts (reduces the standard deviation) by 3.2 percent (0.012 / 0.377). The implication of all of this is that it is important that further analysis account for the pronounced seasonality in the data.

III. Simple statistical tests

We now turn to the key question of this paper: Is the spread (standard deviation) of inflation forecasts by survey participants lower in countries that adopt an explicit inflation target? To examine this, we start with some very simple statistics designed to measure whether the spreads $S_{it}(\pi, c)$ and $S_{it}(\pi, n)$ are lower when a country's policymakers employ an inflation target. For clarity, we do this with a series of regressions. In the first one we estimate the following regression for each country separately:

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[1] We also regressed the coefficients against {1, 2, …, 12, 0, …, 0} and {0, …, 0, 1, 2, …, 12}. The slope coefficients (t-statistics) were -0.003 (2.80) and -0.024 (23.1).

(1) $$S_{it}(\pi, c) = \alpha + \sum_{k=2}^{12} \beta_k D_{kt}^m + \gamma Target_{it} + u_{it}^c$$

and similarly for $S_{it}(\pi, n)$. Of course, the variable “Target” is only included if the country was both an inflation and non-inflation targeter during the sample period. In addition, since the available data are different for different countries, each equation is estimated with a different number of observations.

The results are shown in Appendix Tables A1 and A2 and summarised in Table 2. Note, there are only 6 countries that switched regimes during the sample period. For these countries, the average standard deviation for the non-inflation targeting regime is α, for the inflation targeting regime is $\alpha + \gamma$, and γ is the difference. If the dispersion is smaller for inflation targeting regimes, we would expect to see $\gamma < 0$.

Table 2

Comparing the standard deviation of inflation forecasts for countries that switch regime

γ	**Standard deviation of current year's inflation forecast,** $S_{it}(\pi, c)$	**Standard deviation of next year's inflation forecast,** $S_{it}(\pi, n)$
Negative, significant at 5%	Australia, UK	Australia, Canada, UK
Negative, insignificant at 5%	Canada, Spain, Switzerland	
Positive, insignificant at 5%		Switzerland
Positive, significant at 5%	Norway	Norway, Spain

Source: Authors calculations based on country-by-country estimation of equation (1).

These results allow us to conclude that for those countries that adopted inflation targeting during the period 1990 to 2009, the standard deviation of private sector forecasts for CPI inflation sometimes falls and sometimes rises. Specifically, the standard deviation of inflation forecasts falls when Australia, Canada, and United Kingdom adopt inflation targets; but it increases when Norway adopts an inflation target. Spain and Switzerland are somewhere in between, depending on whether you are looking at the dispersion of current year forecasts or next year forecasts.

A next step is to estimate the equations jointly using Zellner's seemingly unrelated regression approach. Since the model is estimated as a set of equation, we need the same number of observations for all countries. There are basically three different start dates that could be used: November 1990, January 1995, or June 1998. Depending on the start date, the number of countries that were both inflation targeters and non-inflation targeters differs, as seen in Table 3. There is clearly a trade-off between sample size, number of countries, and a mix of inflation regimes.

Table 4 summarises the results of estimation using Zellner's seemingly unrelated regression approach, focusing on the sign and significance of the inflation targeting variable for countries that were both inflation targeters and non-targeters. (Additional results are shown in Appendix Table A3.) In general, the results are similar to the single equation regressions in Table 2. Namely, sometimes the coefficient is negative and sometimes it is positive; sometimes it is significant different from zero at standard levels of statistical significance and sometimes it is not.

Table 3

Inflation regime status for various start dates

Start date	Switched	Non-targeter	Targeter only
November 1990 (T = 222, N = 8)	Australia, Canada, UK	France, Germany, Italy, Japan, US	
January 1995 (T = 172, N = 12)	Spain	France, Germany, Italy, Japan, Netherlands, US	Australia, Canada, New Zealand, Sweden, UK
June 1998 (T = 131, N =14)	Norway, Spain, Switzerland	France, Germany, Italy, Japan, Netherlands, US	Australia, Canada, New Zealand, Sweden, UK

Source: Table 1.

Table 4

Comparing the standard deviation for countries that have been both inflation targeters and non-inflation targeters using SUR.

γ	Standard deviation of current year's inflation forecast, $S_{it}(\pi, c)$	Standard deviation of next year's inflation forecast, $S_{it}(\pi, n)$
Negative, significant at 5%	Australia, UK (1990) Spain (1995) Spain (1998)	Australia, UK (1990)
Negative, insignificant at 5%		Canada (1990) Spain (1998)
Positive, insignificant at 5%	Canada (1990) Switzerland (1998)	
Positive, significant 5%	Norway (1998)	Spain (1995) Norway (1998) Switzerland (1998)

Note: The dates in parentheses denote the start dates for the estimation.
Source: Authors' calculations based on SUR estimation of equation (1).

So far, the determinants of the dispersion of private sector forecasts are deterministic variables: seasonal dummies and an inflation targeting dummy. There are no economic variables that may affect the dispersion of forecasts. For example, the dispersion of inflation forecasts may be greater when overall macroeconomic variability is greater. While we do not have a country-specific measure of macroeconomic variability, we use the dispersion of private sector forecasts of GDP growth. To avoid any simultaneity concerns, we actually use the lagged value of the spread of GDP forecasts, $S_{i,t-1}(y,c)$ and $S_{i,t-1}(y,n)$. In addition, we allow the current inflation spread to depend on its own lagged value, $S_{i,t-1}(\pi,c)$ and $S_{i,t-1}(\pi,n)$. That is, we estimate a set of equations of the following form:

(2) $$S_{it}(\pi,c)=\alpha_i^c+\sum_{k=2}^{12}\beta_k^c D_{kt}^m+\gamma^c Target_{it}+\delta^c year_t+\rho_i^c S_{it-1}(\pi,c)+\theta_i^c S_{it-1}(y,c)+u_{it}^c .$$

Since the model is estimated as a set of equations, a *Target* variable can only be included for those countries that switched during the same period.

The results of estimating these equations using Zellner's seemingly unrelated regression approach are summarised in Table 5. With 3 different start dates, the countries that were both inflation targeting and non-inflation targeting differ. We report results only for the countries that change regime, so the coefficient of interest on the *Target* can be estimated.

Table 5

Seemingly unrelated regression estimation

January 1990 – April 2009

	Current year spread			**Next year spread**		
	Target	$S_{i,t-1}(\pi,c)$	$S_{i,t-1}(y,c)$	**Target**	$S_{i,t-1}(\pi,n)$	$S_{i,t-1}(y,n)$
Australia	-0.038	0.585	0.050	-0.026	0.877	0.031
	[1.5]	[10.5]	[0.8]	[1.2]	[27.8]	[0.7]
Canada	-0.055	0.401	-0.014	0.003	0.703	0.041
	[1.2]	[6.5]	[0.4]	[0.1]	[15.2]	[1.4]
UK	-0.014	0.458	0.050	0.002	0.914	0.034
	[.7]	[9.8]	[0.8]	[0.1]	[45.9]	[0.9]
November 1995 – April 2009						
Spain	0.198	-0.015	0.298	-0.059	0.099	0.019
	[11.0]	[1.5]	[4.2]	[0.7]	[2.81]	[1.2]
June 1998 – April 2009						
Norway	0.234	0.067	0.499	0.124	0.106	0.059
	[7.2]	[3.8]	[7.2]	[2.2]	[2.7]	[3.2]
Spain	0.202	-0.022	0.459	-0.023	0.118	-0.005
	[13.0]	[1.2]	[6.2]	[1.4]	[4.2]	[0.2]
Switzerland	0.255	0.000	0.588	-0.011	0.134	0.008
	[12.7]	[0.0]	[8.6]	[0.2]	[4.3]	[0.6]

Note: Numbers in brackets are asymptotic t-ratios.
Source: Authors' calculations based on equation (2).

A few conclusions can be drawn from Tables 5. First, the coefficient on Target is generally insignificant. However, the coefficient is positive and significant for Spain, Norway, and Switzerland in the equation for $S_{it}(\pi,c)$. Second, the coefficient on the lagged value of the spread for the inflation forecast is positive and significant; the coefficient is larger for the next year's inflation forecast than for the current year's inflation forecast. And third, in looking at the full set of regressions not reported in the table, the coefficient on the lagged value of *S(y,.)* is generally insignificant. When the coefficient is significant, it is mostly positive, although there are a couple cases when the coefficient is negative and significant.

While these results are interesting, they fail to utilise information from the countries that either never adopted an inflation target or did so prior to the beginning of our sample. We now turn to a more sophisticated analysis designed to account for seasonality, control for general macroeconomic volatility, and employ all of the data we have available.

IV. Panel regressions

The various shortcomings mentioned at the end of the previous section can be addressed by estimating a set of equation using a panel regression approach. By estimating various regressions using both fixed effects and random effects, we show that there is little evidence that inflation targeting countries have a smaller dispersion of private sector inflation forecasts.

We start by estimating some fixed effects models given the general applicability of this approach. To capture the fact that volatility may have changed over time reflecting the so-called "Great Moderation" and then the more recent financial crisis, several dummy variables are included in the equation. Specifically, year dummy variables are included ($Year_{1989}$, …, $Year_{2006}$), where, for example, $Year_{1989}$ = 1 for 1989. We also define $Year_{2007}$ = 1 for January 2007 – August 2007, $Crisis_1$ = 1 for September 2007 – September 2008, and $Crisis_2$ = 1 for October 2008 – April 2009. These last two variables reflect the fact that the financial crisis is generally thought to have started in August 2007 and then intensified in September 2008. Since the survey respondents presumably did not recognise the start and intensification of the crisis until the next month, the *Crisis* variables are dated one month after the start and intensification. The results from estimating the following equation (or some variant of it) are reported in Table 6 and the time-varying constants are shown in Graph 4[2]:

(3)
$$S_{it}(\pi,c) = \alpha^c + u_i^c + \sum_{k=2}^{12} \beta_{ik}^c D_{kt}^m + \sum_{t=1991}^{2007} \delta_t Year_t + \kappa_1 Crisis_{1it} + \kappa_2 Crisis_{3it}$$
$$+ \gamma_i^c Target_{it} + \rho^c S_{1t-1}(\pi,n) + \theta^c S_{1t-1}(y,n) + \varepsilon_{it}^c$$

where the constant α can be thought of as a mixture of the average level in January, the base month, plus the average impact of October 1989 to December 1990, while the *u* is the country-specific fixed effect.

In all cases, the coefficient on *Target* is negative suggesting that countries with an inflation target have a smaller standard deviation of private sector inflation forecasts. However, the coefficient is significant only for the bare-bones regression (excluding lagged values of $S(\pi,.)$ and $S(y,.)$. In addition, the coefficient on lagged values of $S(\pi,.)$ and $S(y,.)$ are positive and significant. The estimate of ρ is 0.49 for regressions using $S_{it}(\pi,c)$ and 0.74 when looking at $S_{it}(\pi,n)$. This suggests that the persistence of the spread is less for current year forecasts than for next year forecasts, but still sizeable. A larger coefficient on the lagged spread for next year forecasts than for current year forecasts might suggest that incoming monthly data plays a smaller role for next year forecasts than for current year forecasts. In addition, to the extent that lagged values of $S(y,)$ capture overall macroeconomic uncertainty, the positive and significant coefficients suggest that the standard deviation of inflation forecasts depend on macroeconomic uncertainty. Chart 4 shows the effect of time and the financial crisis on the standard deviation of private sector forecasts of current and next year inflation made in

2 Specifically, the value plotted for the years 1991 – 2007 are = $\alpha + \delta_t$ for 2008 = $\alpha + \kappa_1$, and for 2009 = $\alpha + \kappa_2$ reflecting the fact that these are the "constant" terms for January of the specific year.

January. In general, the standard deviation declined from 1991 to 1999, was stable through 2006, and then rose significantly in 2008 and 2009. This suggests that the "Great Moderation" did lead to a reduction in the dispersion of private sector forecasts of inflation through the first part of the sample period and that the financial crisis led to an increase in the dispersion. Interestingly, the coefficient on *(α + δ)* in 2009 is about the same as in the early 1990 for $S(\pi, n)$ but higher for $S(\pi, c)$.

Table 6

Panel estimation using fixed effects

	Standard deviation of current year's inflation forecasts, $S_{it}(\pi, c)$			**Standard deviation of next year's inflation forecasts,** $S_{it}(\pi, n)$		
Inflation Target (γ)	-0.023 [3.0]	-0.011 [1.4]	-0.009 [1.3]	-0.021 [2.0]	-0.017 [1.6]	-0.009 [1.3]
Lagged $S_{it}(\pi,)$ (ρ)			0.488 [29.1]			0.737 [57.3]
Lagged $S_{it}(y,)$ (θ)		0.142 [9.1]	0.036 [2.6]		0.103 [6.1]	0.015 [1.3]
R^2 overall	0.36	0.44	0.62	0.24	0.29	0.75
σ_ε	0.084	0.083	0.073	0.119	0.118	0.180
N	2895	2880	2880	2895	2880	2880

Note: Asymptotic t-ratios are in brackets. Coefficients on monthly and yearly dummy variables are not shown (but were included).
Source: Authors' calculations based on equation (3).

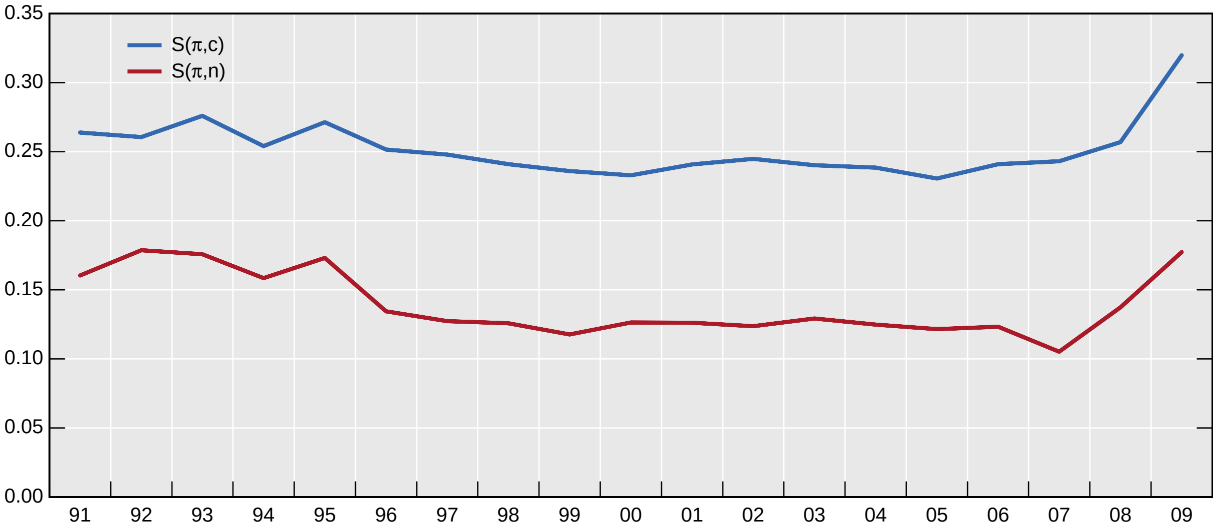
While the coefficient on *Target* is negative for both $S(\pi, c)$ and $S(\pi, n)$, the magnitude is economically small. There are several ways to see this. First, look back at Figures 1 and 2 and note that the range of $S(\pi, c)$ is roughly 0.25 while $S(\pi, n)$ averages closer to 0.40. This means that the impact of inflation targeting is to reduce the standard deviation of inflation forecasts by between one part in 25 and one part in 40 on average.

A second way to see that γ is economically small is compare it to the impact of the financial crisis. To do this, we consider what would happen to $S(\pi, c)$ if a country were to adopt an inflation target in February 2007 versus adopting it in February 2009. Taking the case of the United States, we use equation (2) to calculate the predicted value of *S(π, c)* first with the *Target* = 0 (no inflation target) and then with the *Target* = 1 (assuming the U.S. had an inflation target). The result is that the predicted dispersion in survey inflation expectations would fall from 0.269 to 0.260 in February 2007, but from 0.663 to 0.654 in February 2009. In other words, while dispersion would be less, the impact of financial turmoil on the dispersion is much larger – 0.394 versus 0.009, or 43 times larger.

We next considered estimating the panel regression using a random effects estimator. The fixed effects (FE) model specifies the country specific effect as a constant, whereas the random effects (RE) model specifies the country specific effect as a random variable that is uncorrelated with the regressors. Breusch and Pagan (1980) developed a Lagrange multiplier test for $\sigma_u^2 = 0$; the p-value is reported in Table 9 in the row labeled "$\sigma_u^2 = 0$ p-value." If the orthogonality assumption is true, then the random effects model is more efficient because it uses the assumption that u_i is uncorrelated with the regressors. Of course, if this assumption is false, then the random effects model is inconsistent. We can then use a Hausman test of the extra orthogonality condition imposed by the random effects estimator. The idea of the Hausman test is simple: If the regressors are uncorrelated with u_i, the fixed effects estimator is consistent but inefficient and the random effects estimator is consistent and efficient. However, if the regressors are correlated with u_i, the fixed effects estimator is consistent but the random effects estimator is inconsistent. Table 9 estimates the same models and includes the p-value from the Hausman test in the row labelled "Hausman".[3]

Graph 4

Time-varying constant term



Source: Author's calculations.

The results from using random effects to estimate the model are mixed. The coefficient on *Target* is sometimes negative and insignificant and other times it is positive and significant. Not surprisingly, one can always reject the hypothesis that $\sigma_u^2 = 0$. Unfortunately, one can often reject the hypothesis that the orthogonality condition holds. In general, it appears as though we cannot reject the hypothesis that orthogonality condition holds for the stripped-down model (which includes only *Target* and year and month dummies) but we can generally reject the orthogonality condition when we include lagged values of $S(\pi,.)$ and $S(y,.)$. Interestingly, in the bare-bones model (which fails to reject the Hausman test) the coefficient on *Target* is negative but insignificant.

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[3] Including a lagged dependent variable in a fixed effects model creates a large-sample bias in the estimate of the coefficient on lagged dependent variable. Since we are not really interested in this estimate, the concern is somewhat mitigated. In addition, in a simple model Nickell (1981) shows that for large values of T, the limit of $(\hat{\rho} - \rho)$ as $N \to \infty$ is approximately –(1+ρ)/(T-1). With ρ = 0.5 (0.8) and T = 235, the bias will be -0.0064 (-0.0077).

Table 8

Panel estimation using random effects

	Standard deviation of current year's inflation forecasts, $S_{it}(\pi, c)$			Standard deviation of next year's inflation forecasts, $S_{it}(\pi, n)$		
Inflation Target (γ)	-0.004 [0.6]	0.013 [2.1]	0.022 [7.2]	-0.012 [1.2]	-0.007 [0.7]	0.014 [4.2]
Lagged $S_{it}(\pi,)$ (ρ)			0.542 [33.2]			0.785 [66.3]
Lagged $S_{it}(y,)$ (θ)		0.168 [11.2]	0.082 [7.0]		0.110 [6.5]	0.040 [4.2]
R^2 overall	0.39	0.49	0.65	0.25	0.31	0.76
σ_ε	0.084	0.083	0.073	0.119	0.118	0.080
$\sigma_u^2 = 0$ p-value	0.000	0.000	0.000	0.000	0.000	0.000
Hausman p-value	0.276	0.000	0.000	0.961	0.999	0.000
N	2895	2880	2880	2895	2880	2880

Note: Asymptotic t-ratios are in brackets. Coefficients on monthly dummy variables are not shown (but were included).
Source: Author's calculations based on equation (3).

V. Conclusions

Using survey data on inflation expectations drawn from Consensus Forecast, we find little evidence that inflation targeting countries have a smaller dispersion of private sector forecasts of inflation. While for some countries, for some models, and some estimation techniques, we estimate that inflation targeting countries have a smaller dispersion of private sector inflation forecasts, for other countries, other models, and other estimation techniques, we find that they do not.

Returning to the question that motivated this analysis – "Does increased transparency lead to lower dispersion in private forecasts?" – the answer appears to be no. This suggests to us that the Morris and Shin argument that increased transparency could be destabilising is of little practical concern to policymakers. Of course, since the survey we use only reports results of forecasts for inflation in the current year and the next year, our results are unable to shine light on whether the distribution of private sector forecast of long-run inflation is lower. However, even if the inflation target is for the medium-run, one would expect that the dispersion of inflation forecasts for "next year" may still be somewhat smaller than otherwise.

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Appendixes

Appendix Table A1

Average standard deviation of current year inflation forecasts

	α—January			
	No inflation targeting	**Inflation targeting**	**Difference**	**Number of observations**
Australia	0.510 [17.7]	0.369 [15.4]	-0.141 [7.17]	215
Canada	0.291 [16.3]	0.290 [22.96]	-0.001 [0.09]	228
Norway	0.264 [7.8]	0.397 [12.7]	0.133 [6.7]	124
Spain	0.221 [13.1]	0.212 [11.6]	-0.009 [0.8]	165
Switzerland	0.307 [13.6]	0.288 [15.6]	-0.019 [1.32]	124
UK	0.458 [14.3]	0.301 [11.3]	-0.157 [7.4]	228
Euro Area	0.147 [11.4]			70
France	0.208 [18.3]			228
Germany	0.207 [25.8]			228
Italy	0.231 [12.2]			228
Japan	0.275 [11.0]			228
Netherlands	0.243 [13.3]			165
US	0.310 [20.0]			228
New Zealand		0.436 [12.4]		166
Sweden		0.324 [14.8]		165

Appendix Table A2

Average standard deviation of next year inflation forecasts

	α—January			
	No inflation targeting	**Inflation targeting**	**Difference**	**No inflation targeting**
Australia	0.780 [14.2]	0.501 [10.7]	-0.296 [7.7]	215
Canada	0.437 [14.2]	0.372 [17.2]	-0.064 [2.6]	228
Norway	0.224 [6.3]	0.334 [10.2]	0.110 [5.3]	124
Spain	0.283 [7.3]	0.399 [9.4]	0.117 [4.6]	165
Switzerland	0.271 [10.0]	0.299 [10.2]	0.029 [1.6]	124
UK	0.764 [11.6]	0.495 [9.0]	-0.269 [6.2]	228
Euro Area	0.203 [10.2]			70
France	0.237 [17.3]			228
Germany	0.308 [17.3]			228
Italy	0.309 [8.7]			228
Japan	0.403 [11.9]			228
Netherlands	0.334 [10.3]			165
US	0.439 [21.1]			228
New Zealand		0.431 [12.4]		166
Sweden		0.321 [11.3]		165

Appendix Table A3

Coefficient on inflation targeting dummy variable

		November 1990 – April 2009	January 1995 – April 2009	June 1998 – April 2009
Australia	S(π, c)	-0.136 [6.9]		
	S(π, n)	-0.195 [6.3]		
Canada	S(π, c)	0.038 [1.1]		
	S(π, n)	-0.010 [0.02]		
UK	S(π, c)	-0.082 [4.5]		
	S(π, n)	-0.085 [2.7]		
Spain	S(π, c)		-0.021 [2.3]	-0.037 [2.2]
	S(π, n)		0.080 [4.9]	-0.002 [0.1]
Norway	S(π, c)			0.127 [7.7]
	S(π, n)			0.106 [6.3]
Switzerland	S(π, c)			0.003 [0.3]
	S(π, n)			0.032 [2.4]